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HYSTERESIS IN UNEMPLOYMENT: EVIDENCE FROM NORWEGIAN COUNTIES

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Hysteresis in Unemployment: Evidence from Norwegian Counties*

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Abstract

The paper tests hysteresis effects in unemployment using aggregate and panel data for Norway. While tests using aggregate or county–specific time series do not reject the null of unit root hysteresis, the panel tests firmly reject the null. When a one–time structural break is incorporated, the unit root hypothesis is rejected (or nearly so) in most counties. All results reveal a high degree of unemployment persistence, but the speed of adjustment is much higher when we allow for a change in mean.

Key words: Hysteresis, panel unit root tests, structural break.

JEL classifications: C22, C23, J64.

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1 Introduction

This paper investigates empirically the possibility of unemployment hysteresis in Norway, applying panel data for 19 counties for the time period 1961 to 1998. The term hysteresis is used to describe a situation in which transitory shocks to the unemployment rate have permanent effects. The tests reported in this paper are carried out within the framework of a linear dynamic model, in which hysteresis requires a unit root in the unemployment process. While tests based on aggregate time series and time series data for individual counties do not reject the null hypothesis of unit root hysteresis the panel unit root tests firmly reject the null hypothesis. We also report results for unit root tests in the presence of structural change. When we allow for a one–time change in mean, the unit root hypothesis is rejected for most counties.

Previous studies of unit root hysteresis using aggregate time series data include Barro (1988), Blanchard and Summers (1986), Jaeger and Parkinson (1994),[1] Mitchell (1993), Leslie *et al.* (1995) and Røed (1996). Mitchell (1993) employs alternative test procedures to quarterly unemployment rates for 15 industrialized countries, covering the period from the mid 1960s to 1991. For some countries he also uses annual time series covering more than 100 years. His tests consistently fail to reject unit roots even when structural break dummies are included in the regressions. Similar results based on quarterly time series for 16 OECD countries for the period 1970–94 can be found in Røed (1996). The null of a unit root is only rejected for the USA. Leslie *et al.* (1995) report results for various unit root tests (including deterministic trend terms) using yearly unemployment rates for 23 countries. For most countries the sample period is from 1948 to 1992. Almost all tests fail to reject the presence of unit roots, with some exceptions for USA, Israel and New Zealand. Papell *et al.* (2000) report Augmented Dickey–Fuller (ADF) tests (with and without a deterministic trend) using annual unemployment series from 1955 to 1997 for sixteen OECD countries. In all cases, the null of a unit root can not be rejected as long as structural breaks are not incorporated.

There are several problems with these tests. First, testing for hysteresis within the framework of a linear dynamic model may be considered to constitute a rather extreme special case since the root has to be exactly equal to one in order to produce hysteresis. It may be argued that hysteresis is ultimately a non–linear phenomenon, associated with the possibility of multiple

[1] Jaeger and Parkinson (1994) report results for standard unit root tests. They also propose an unobserved components model where unemployment is decomposed into a natural rate component and a cyclical component.

stable unemployment equilibria.[2]

The second problem concerns the power of standard unit root tests. It is by now generally accepted that the commonly used unit root tests like the ADF and the Phillips–Perron tests lack power in distinguishing the unit root null from stationary alternatives, in particular when the root is close to unity. Using panel data unit root tests is one way of increasing the power of tests based on a single time series. Intuitively, combining information from the time series dimension with that obtained from the cross-sectional will increase the sample size and therefore make inference more precise. Also, while test statistics based on time series information have been shown to have complicated limiting distributions, the corresponding statistics based on non-stationary panel data have been shown to converge to Gaussian distributions.[3] In this paper we mainly apply the tests proposed by Im, Pesaran and Shin (1997), IPS, to panel data for county–specific Norwegian unemployment rates. These tests allow for complete parameter heterogeneity across counties. Results based on the Levin and Lin (1992), LL, test where only deterministic terms vary across counties are briefly considered.

Previous studies using panel data for unemployment rates include Song and Wu (1997, 1998) who employ the LL test. While Song and Wu (1997) make use of panel data for US states, Song and Wu (1998) test the null of a unit root in the unemployment process by pooling data for OECD countries. In both studies, the panel unit root tests reject the null hypothesis that the unemployment rate is a random walk process.

Finally, it has become well known that nonrejection of the unit root hypothesis may be caused by mis–specification of the deterministic components. Using tests developed by Perron and Vogelsang (1992) for nontrending data, we test for a unit root while allowing for a one–time break in the mean, where the location of the breakpoint is unknown a priori. When a one–time structural break is incorporated, the results in Papell *et al.* (2000) indicate that the unit root hypothesis is rejected for most of the sixteen OECD counties included in their study.

The rest of the paper is organized as follows. Section 2 reports some

[2] Theoretical models producing multiple unemployment equilibria are formulated in Diamond (1982), Cooper and John (1988), Drazen (1987), Manning (1990, 1992), Saint-Paul (1995) and Moene *et al.* (1997). Tests for multiple unemployment equilibria within a non-linear dynamic framwork can be found in Bianchi and Zoega (1997, 1998) using Markov Switching Regressions (MSR) and in Akram (1998, 1999) using both MSR and Smooth Transition Regressions.

[3] We refer to the special issue of Oxford Bulletin of Economics and Statistics, Vol 61, November 1999, for a colletion of papers on unit root testing and cointegration using panel data. See in particular Banerjee (1999) for an overview, and Maddala and Wu (1999) for a comparative study.

preliminary facts about unemployment in Norway. In Section 3 we report unit root tests based on aggregate unemployment, county specific regressions and the panel unit root tests. Section 4 concludes.

2 Data and some facts

The unemployment rates used in this study are defined as the number of openly or registered unemployed persons divided by the labour force. These numbers do not include participants joining labour market programs. For aggregate unemployment we have consistent time series from 1948 to 1998 while the sample period for county specific unemployment rates is 1961 to 1998.

Figure 1 graphs aggregate open unemployment for the time period 1948–1998 while the distribution in aggregate unemployment over time is graphed in Figure 2. With an exception for the late 1950s, unemployment was below 2 per cent until the early 1980s and increased to 3.3 per cent in 1984. The unemployment rate was reduced to 1.5 per cent in 1986 for thereafter to increase to 5.5 per cent in 1993. From 1993 the rate of open unemployment gradually declined to 2.4% in 1998, and has been more or less constant thereafter. A remarkable feature with the Norwegian unemployment experience – as compared with most European countries – is therefore that unemployment has been reduced relatively soon after an unfavourable shock. Moreover, unemployment seems to fluctuate more during the last part of our sample period. Figure 2 shows that the aggregate rate of unemployment clusters around 1.5%. The distribution is skewed to the right, but the Figure does not display a bimodal density (or there is only a weak tendency of a second hump).[4]

Figure 1–2 about here

Figures 3 to 6 give a description of the regional distribution in unemployment. Figure 3 shows that county unemployment rates largely move in tandem. When we regressed county–specific unemployment changes against changes in aggregate unemployment we obtained an average $\overline{R}^2$ equal to 0.78. Hence, shocks to the unemployment rate is largely symmetric across counties. Figure 4 shows that the maximum value of regional unemployment approximates 3% until the early 1980s while the minimum value is close to

[4] Using quarterly data 1972:1–1998:1, Akram (1999) reports results indicating a bimodal density with modes centered at around 2 and 5 per cent, respectively.

zero. There is a tendency that the minimum value increases more than the maximum value during slumps.

To take a closer look at the cross–county dispersion we calculate variances of relative unemployment, $Var\,(U_{it}/U_t)$, which can be interpreted as a measure of regional mismatch (Layard *et al.* 1991, Ch. 6). From Figure 5 we first note that the variance of relative unemployment is very high during the 1960s and the first part of the 1970s as compared to corresponding figures for other countries reported in Layard *et al.* (1991, pp. 294–5). We also see that regional mismatch decreases during the last part of the sample period. Comparing Figure 1 and 5 we see that regional mismatch is reduced during downturns and widen during booms. Although the decrease in mismatch can not be found in any other country investigated by Layard *et al.* (1991), this result is highly consistent with the finding in Dyrstad and Johansen (2000) using data for Norwegian municipalities for the time period 1970–88. When interpreting these results we should have in mind that we use data for open unemployment. Using data for total unemployment may modify the result of decreasing variances as labour market programs have been targeted towards high unemployment regions.

Figure 3 – 6 about here

Figure 6 investigates the possibility of convergence across counties. We first calculate the average unemployment rate for each county for the 1960s and the 1990s. Based on these numbers we plot the changes in average unemployment against the mean for the 1960s. Figure 6 shows that average unemployment increased in all counties, but increased most in counties with low initial levels.[5] This result is also in accordance with the findings in Dyrstad and Johansen (2000).

3 Tests for unit roots

3.1 Aggregate unemployment

Before we turn to the panel data tests of unemployment hysteresis we present results based on aggregate time series data for the time period 1948 – 1998.

[5] Regressing the changes against the initial levels gave (t-statistics in parentheses)

$$\overline{U90}_i - \overline{U60}_i = \underset{(12.29)}{3.98} - \underset{(-2.70)}{0.60}\,\overline{U60}_i, \text{ with } \overline{R}^2 = 0.26,$$

where $\overline{U60}_i$ and $\overline{U90}_i$ are the average unemployment rate in county i in the 1960s and 1990s, respectively.

To test the null hypothesis of random walk, we run regressions like

$$\Delta U_t = \alpha U_{t-1} + \sum_{j=1}^{P} \beta_j \Delta U_{t-j} + \eta + \gamma t + \varepsilon_t, \tag{1}$$

where $\alpha = 0$ under the null hypothesis and negatively signed under the alternative. Results based on equation (1) with $\gamma = 0$ are given by (t-statistics in parentheses)

$$\Delta U_t = \underset{(-2.43)}{-0.103} U_{t-1} + \underset{(4.63)}{0.584} \Delta U_{t-1} + \underset{(2.07)}{0.209} \tag{2}$$

$$\Delta U_t = \underset{(-1.52)}{-0.067} U_{t-1} + \underset{(5.28)}{0.718} \Delta U_{t-1} - \underset{(-2.23)}{0.339} \Delta U_{t-2} + \underset{(1.47)}{0.150}, \tag{3}$$

for $P = 1$ and 2, respectively. The Dickey-Fuller "t"–statistics are –2.43 and –1.52 and well above the 5% critical value equal to –2.92. Hence the null of a root equal to unity can not be rejected.

When equations (2) and (3) are expanded with a linear deterministic trend we obtain the following results

$$\Delta U_t = \underset{(-3.01)}{-0.185} U_{t-1} + \underset{(4.96)}{0.619} \Delta U_{t-1} + \underset{(1.80)}{0.011} t + \underset{(0.67)}{0.081}, \tag{4}$$

$$\Delta U_t = \underset{(-2.13)}{-0.140} U_{t-1} + \underset{(5.43)}{0.730} \Delta U_{t-1} - \underset{(-2.13)}{0.287} \Delta U_{t-2} + \underset{(1.48)}{0.009} t + \underset{(0.35)}{0.044} \tag{5}$$

where the Dickey–Fuller statistics are now –3.01 and –2.13, respectively, compared with the 5% critical value equal to –3.50. Again we can not reject the null hypothesis of a unit root in the Norwegian unemployment process.[6] In all equations, we note that the estimated speed of adjustment parameter, $|\alpha|$, is small which implies very sluggish adjustments. As is well known, the OLS estimates of $|\alpha|$ are biased upwards which means that the true values are even closer to zero.[7] Finally, we note a high degree of autocorrelation in unemployment changes.

3.2 County specific regressions

Turning to the time series properties of regional unemployment, we first estimate county specific regressions like

$$\Delta U_{it} = \alpha_i U_{it-1} + \beta_i \Delta U_{it-1} + \eta_i + \varepsilon_{it}, \quad i = 1, 2, ..., 19, \tag{6}$$

[6] This case is rather extreme as $\alpha = 0$ implies a linear trend in unemployment changes and a quadratic trend in the level.

[7] This comment also applies to the estimates presented below.

where all parameters are unrestricted in the cross section dimension.[8]

Results based on county–specific regressions are reported in Table 1. We first note that the estimates of α_i are consistently negative across counties with a mean of –0.13. Again, the results imply low speed of adjustment – and a high degree of unemployment persistence – in all counties. The cross–county variation in the estimated speed of adjustment parameter is moderate as $|\widehat{\alpha}_i|$ ranges from 0.08 to 0.24. In most cases the estimates are slightly above 0.1. Most importantly, using data for individual counties the null hypothesis of a unit root is only rejected for one county using 10% critical values. Finally, all county–specific regressions reveal that unemployment changes are highly autocorrelated.

Table 1 about here

3.3 Panel unit root tests

Turning to the panel unit root tests we first apply the *"t–bar"* and *"LM–bar"* statistics proposed by IPS. The null hypothesis is that unemployment is a unit root process in all counties ($\alpha_i = 0 \ \forall \ i$), while the alternative is that at least one of the unemployment series is stationary.[9] The IPS *t-bar* statistic is simply defined as the average of the individual Dickey–Fuller t statistics, that is

$$\bar{t}_{NT} = \frac{1}{N} \sum_{i=1}^{N} t_{iT}, \tag{7}$$

where t_{iT} is the usual $t-$statistic, testing the null hypothesis that $\alpha_i = 0$ against the alternative that at least one of the county specific unemployment rates are stationary. Exact sample critical values of $\bar{t}_{NT}$ can be found in IPS, Table 4.

Under the null hypothesis, each of the $t-$statistics can be regarded as a random draw from a Dickey–Fuller distribution with expectation $E\left[t_{iT}|\alpha_i = 0\right]$ and variance $V\left[t_{iT}|\alpha_i = 0\right]$. IPS further propose the use of the standardized *t–bar* statistic defined by

$$\Gamma_{\bar{t}} = \frac{\sqrt{N}\left\{\bar{t}_{NT} - E\left[t_{iT}|\alpha_i = 0\right]\right\}}{\sqrt{V\left[t_{iT}|\alpha_i = 0\right]}}. \tag{8}$$

[8] We also estimated equation (6) with further lag augmentation but one lagged differences maximises the Akaike information criterion in all counties.

[9] The IPS–tests assume that the error terms are independent across counties.

The means, $E\left[t_{iT}|\alpha_i = 0\right]$, and variances, $V\left[t_{iT}|\alpha_i = 0\right]$ are tabulated in IPS, Table 2, based on stochastic simulations. IPS conjecture that the standardized *t–bar* statistic converges weakly to a standard normal variate under the null hypothesis as both N and $T \to \infty$ and $N/T \to k$, where k is a finite positive constant.

The *LM–bar* statistic is defined as the average of the individual LM statistics,

$$\overline{LM}_{NT} = \frac{1}{N}\sum_{i=1}^{N} LM_{iT}, \tag{9}$$

where LM_{iT} is the individual unit root LM statistic for testing the hypothesis that $\alpha_i = 0$ against the alternative that $\alpha_i < 0$. Exact sample critical values of $\overline{LM}_{NT}$ are tabulated in IPS, Table 3. Again, IPS propose to base the test on the standardized *LM–bar* statistic

$$\Gamma_{\overline{LM}} = \frac{\sqrt{N}\left\{\overline{LM}_{NT} - E\left(LM_{iT}\left|\alpha_i = 0\right.\right)\right\}}{\sqrt{V\left(LM_{iT}\left|\alpha_i = 0\right.\right)}} \tag{10}$$

where the means $E\left(LM_{iT}\left|\alpha_i = 0\right.\right)$ and variances $V\left(LM_{iT}\left|\alpha_i = 0\right.\right)$ are obtained from stochastic simulations and tabulated in IPS, Table 1. IPS show that the standardized *LM–bar* statistic $\Gamma_{\overline{LM}}$ converges weakly to a standard normal distribution as N and $T \to \infty$.

Results for the IPS tests for unit root are reported in Table 2. The average value of the $t-$ statistic is –2.09. Using exact sample critical values we reject the null of unit root against the alternative that unemployment is stationary in at least one county at both 5% and 1% levels of significance. The estimated value of the standardized $t - bar$ statistic, defined in equation (6), is –2.74 and well below the 1% critical value for a normal variate. The average value of the LM statistic of 4.12 approximates its exact sample 1% critical value while the standardized $LM - bar$ statistic defined in equation (10) clearly reject the null hypothesis.

Table 2 about here

To investigate the robustness of the results obtained above, we also apply the LL panel unit root test. The test used here is performed by estimating equation (4) with $\alpha_i = \alpha$ and $\beta_i = \beta$ which means that cross–county parameter homogeneity is imposed.[10] The null of unit root, $\alpha = 0$, is tested against

[10] A Wald test of the cross–county restrictions yields $\chi(36) = 27.64$ with a p–value of 0.83.

the alternative that $\alpha < 0$ in *all* counties.[11]

The fixed–effects model is estimated using the within estimator to obtain

$$\Delta U_{it} = -\underset{(-8.72)}{0.121} U_{it-1} + \underset{(12.7)}{0.453} \Delta U_{it-1} + \widehat{\eta}_i, \tag{11}$$

with one lagged difference while expanding with a second lag gives

$$\Delta U_{it} = -\underset{(-7.06)}{0.105} U_{it-1} + \underset{(13.1)}{0.491} \Delta U_{it-1} - \underset{(-3.39)}{0.139} \Delta U_{it-2} + \widehat{\eta}_i. \tag{12}$$

The LL–statistic is –8.72 with one lag and –7.06 with two lagged differences. The 5% critical value approximates –7.0.[12] The null of a unit root is therefore rejected confirming the results based on the IPS tests. Thus, it is tempting to conclude that the inability to reject unit root hysteresis using univariate time series tests reflects the low power against local stationary alternatives. The results reported above are fully in accordance with the evidence reported in Song and Wu (1997, 1998) using data for US states and OECD countries, respectively.[13]

3.4 Unit roots in the presence of structural change

While the results based on panel data are evidence in favour of (a weak version of) the natural–rate hypothesis, the underlying assumption of constant equilibrium unemployment rates may be too restrictive. The equilibrium (or structural) rate of unemployment may well change over time due to permanent supply side shocks. Within theoretical models based on imperfect competition (Layard *et al.* ,1991, 1994; Lindbeck, 1993; and Nickell, 1998, *inter alia*) the equilibrium rate is influenced by factors like the generosity of unemployment benefits, union power, product market competition and the tax wedge.

In this section we investigate the possibility of a one–time change in mean, using data for aggregate as well as county–specific unemployment rates. Testing for unit roots in the presence of structural change has attracted a great deal of interest since the influential paper by Perron (1989). Like Papell *et al.* (2000) we utilize the methods of Perron and Vogelsang (1992) which are

[11] Note that the alternative in the LL test is different from the alternative hypothesis in the IPS tests.

[12] The 5% critical value reported in LL, Table 5 is –7.07 for N=20 and T=25, and –6.32 for N=15, T=25.

[13] Papell *et al.* (2000) report that panel unit root tests "...provide no additional evidence against unit roots among the sixteen unemployment rates", see note 5, p. 310. Detailed results are not reported.

appropriate for nontrending data. We estimate additive outlier (AO) models in which the structural change is assumed to occur instantaneously. The AO model is estimated by the following two equations

$$U_t = \eta + \delta DU_t + e_t \tag{13}$$

where

$$DU_t = \begin{cases} 1 \text{ for } t > TB \\ 0 \text{ otherwise,} \end{cases}$$

$$\Delta\widehat{e}_t = \alpha\widehat{e}_{t-1} + \sum_{j=1}^{P} \beta_j \Delta\widehat{e}_{t-j} + \sum_{j=0}^{P} \omega_j \Delta DU_{t-j} + \xi_t. \tag{14}$$

The location of the breakpoint, TB, is unknown a priori. Equation (13) and (14) are estimated sequentially for each potential break year $TB = P + 2, ..., T - 1$, where T is the sample size. Perron and Vogelsang (1992) consider two procedures for choosing the breakpoint. In the first procedure, the breakpoint is chosen to minimize the $t-$statistic on α. In the second one, the breakpoint is chosen to maximize (or minimize) the $t-$statistic on δ – the coefficient of the change in mean. The second procedure imposes the a priori assumption that the *sign* of the possible change is known (while its location remains unknown a priori). In both cases, however, the procedure for identifying the location of the breakpoint is chosen for statistical, not economic, reasons.[14]

Results based on aggregate unemployment for the time period 1951 to 1998 where the location of the breakpoint is chosen to minimize t_α are given by (t-statistics in parentheses)

$$\begin{aligned} U_t &= \underset{(9.45)}{1.348} + \underset{(8.60)}{2.429} DU_t + \widehat{e}_t & (15) \\ \Delta\widehat{e}_t &= \underset{(-4.83)}{-0.318}\widehat{e}_{t-1} + \underset{(10.38)}{0.668}\Delta\widehat{e}_{t-1} - \underset{(-6.01)}{2.248}\Delta DU_t + \underset{(-2.34)}{1.164}\Delta DU_{t-1}, & (16) \end{aligned}$$

where $DU_t = 1$ for $t > 1985$, zero otherwise. We first note a positive and significant change in mean. Second, the null hypothesis of unit root is rejected at 5% level and nearly rejected at 1% level as the corresponding critical values are –4.25 and –4.95, respectively. Third, the estimated speed of adjustment

[14] The dummy variable ΔDU_{it} is included to allow for a change in mean under the null, and the dummy variables ΔDU_{it-j} are included to ensure that the $t-$statistic on α in equation (14) has the same asymptotic distribution as in the innovative outlier model and is invariant to the value of P, see Perron and Vogelsang (1992) for further details.

coefficient is much higher in the AO model as compared to previous results based on aggregate unemployment.

Results for the AO model using county–specific regressions where the location of the breakpoint is chosen to minimize t_α are reported in Table 3. Results based on the alternative procedure are very similar and available from the author upon request. For all counties, both procedures identify a positive and significant change in mean. The location of the breakpoint differs across counties. While the results in Table 3 indicate break years either in the late 1970s or in the middle of the 1980s, the alternative procedure generally identify break years approximately two years later. Different location of the breakpoint across counties may – at least partly – reflect different industry mix. There is a clear tendency that unemployment increased first in counties dominated by manufacturing industry and some years later in counties dominated by service industries (due to an unfavourable demand shock). We also note that the estimated coefficient of the break dummy is in general larger for counties with initially low mean unemployment (low estimated value of η).

Table 3 about here

Turning to the unit root test, the results in Table 3 reject the null hypothesis at a significance level of 10% for twelve counties. Among these the null is rejected at 5% (1%) for four (two) counties. Based on the alternative procedure for choosing the break year, the null of a unit root is rejected at 10% level of significance in ten countries. Among these the null is rejected in five (one) cases at 5% (1%) level of significance.

The estimated speed of adjustment coefficient is much higher in the additive outlier model as compared with the model with no break. The mean value of the estimates increases from 0.130 (no change in mean) to 0.425 based on the results in Table 3.[15]

Finally, we note that the average of the estimated speed of adjustment coefficient based on county–specific regressions is higher than the corresponding estimate using aggregate unemployment. This result may reflect an equilibrating mechanism that works through migration from high to low unemployment counties.

The average value of t_α, derived from the results in Tables 3, is –4.08. This estimate is in between the 5 and 10% critical values for *univariate* time series. Since all panel unit root tests proposed in the literature are more powerful than their univariate counterparts, we suggest that a test that utilize the

[15] Note that the OLS estimates of α are biased also in the AO model.

cross–sectional variation would firmly reject the unit root null in the additive outlier model.

4 Concluding comments

The paper investigates the time series properties of unemployment in Norway using both aggregate and county–specific data. While tests based on aggregate or county–specific time series do not reject the null hypothesis of a unit root in the unemployment process, the panel unit root tests firmly reject the null hypothesis of unemployment hysteresis. Our findings are consistent with those in Song and Wu (1997, 1998) who report evidence in support of the weak version of the natural–rate hypothesis. These results casts some doubt on previous time series studies that support the unit root hysteresis hypothesis.

Secondly, when a one–time change in mean is incorporated, the unit root hypothesis is rejected for aggregate unemployment. Using county–specific unemployment rates, the null hypothesis is rejected (or nearly so) in most cases. Using data for aggregate as well as county–specific unemployment, we identify a positive and significant change in mean either in the late 1970s or in the middle of the 1980s. The positive change in mean may be interpreted as evidence in favour of increased equilibrium rates of unemployment.

Finally, all results reveal a high degree of unemployment persistence, although the estimated speed of adjustment coefficients are much higher based on the additive outlier model. In the short- and intermediate run, the practical implication of rejecting the null hypothesis of unit root hysteresis against the alternative of highly persistent series may not be very important. But in the long run, unemployment being a random walk process has far more extreme implications as compared with the alternative.

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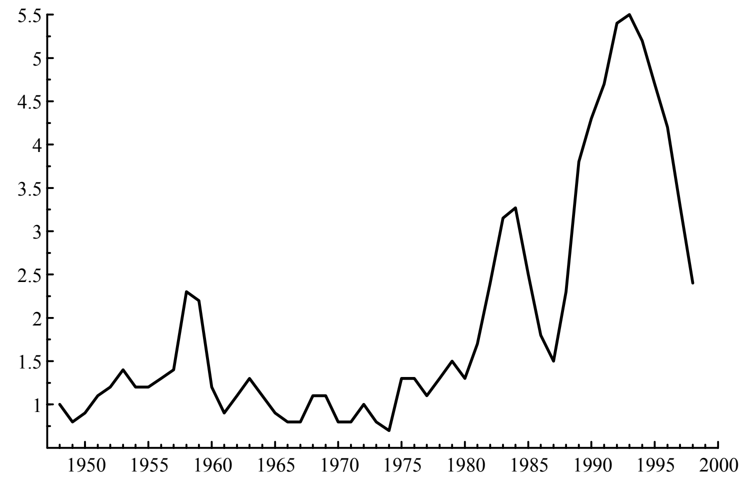

Figure 1: Aggregate unemployment 1948 – 1998, per cent.

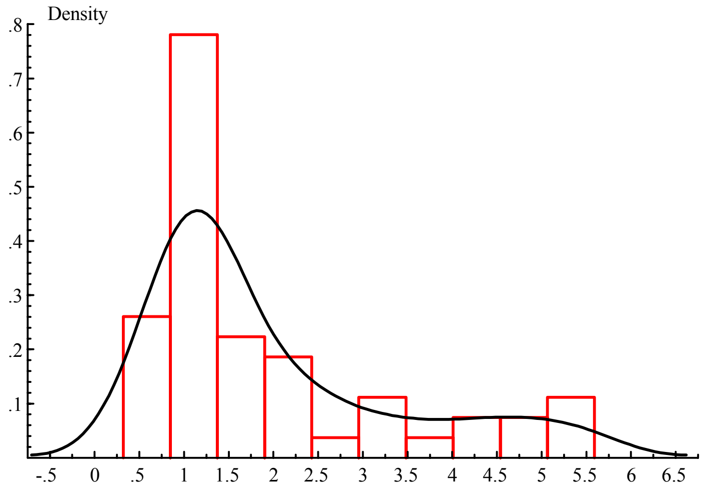

Figure 2: Distribution of aggregate unemployment 1948–1998.

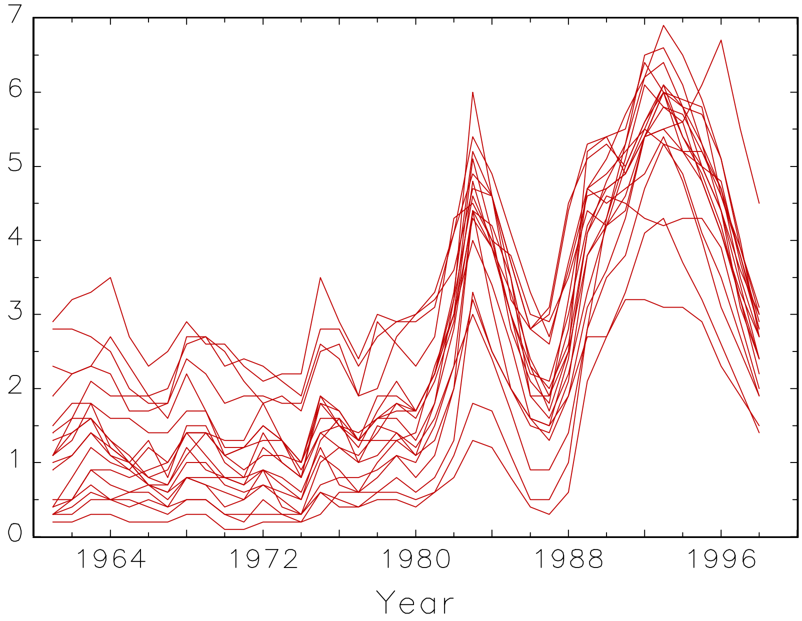

Figure 3: County Unemployment Rates 1961–1998, Per Cent.

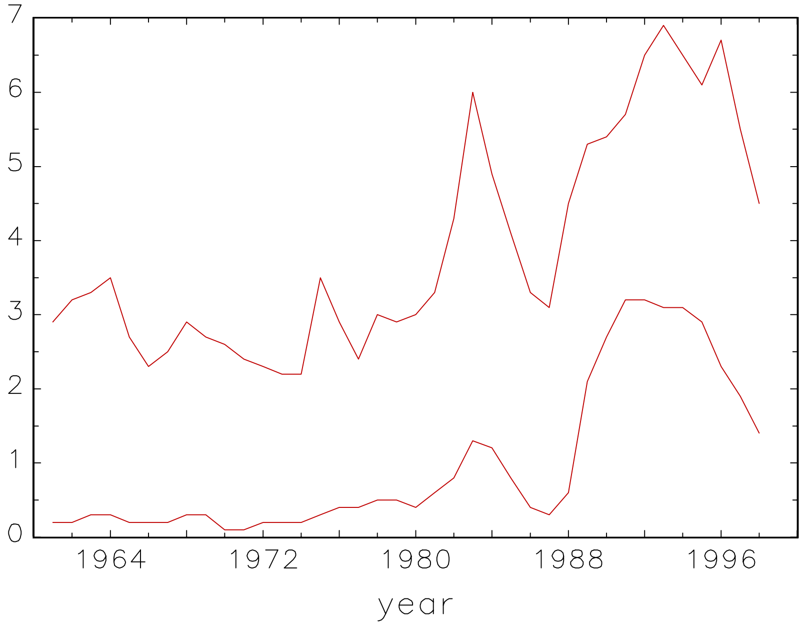

Figure 4: Max and Min Values of County Unemployment Rates, Per Cent.

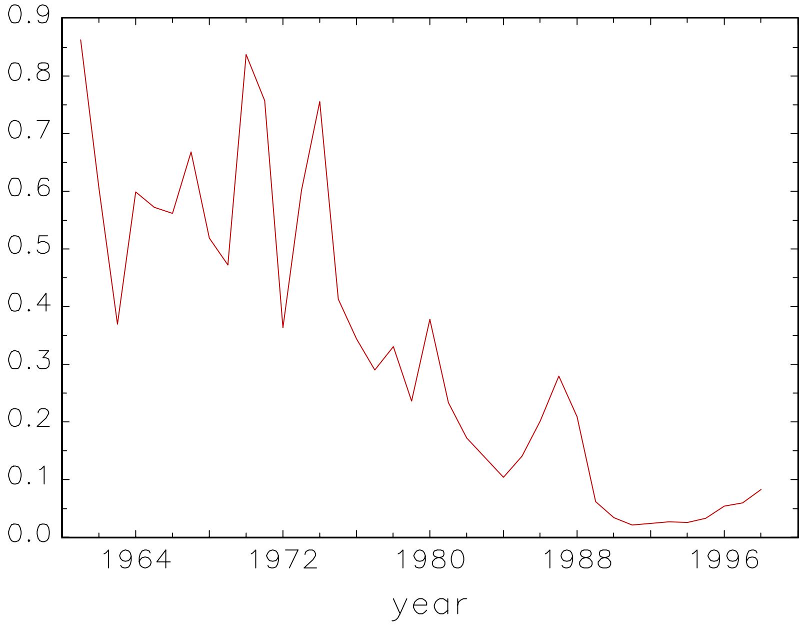

Figure 5: Variances of Relative Unemployment

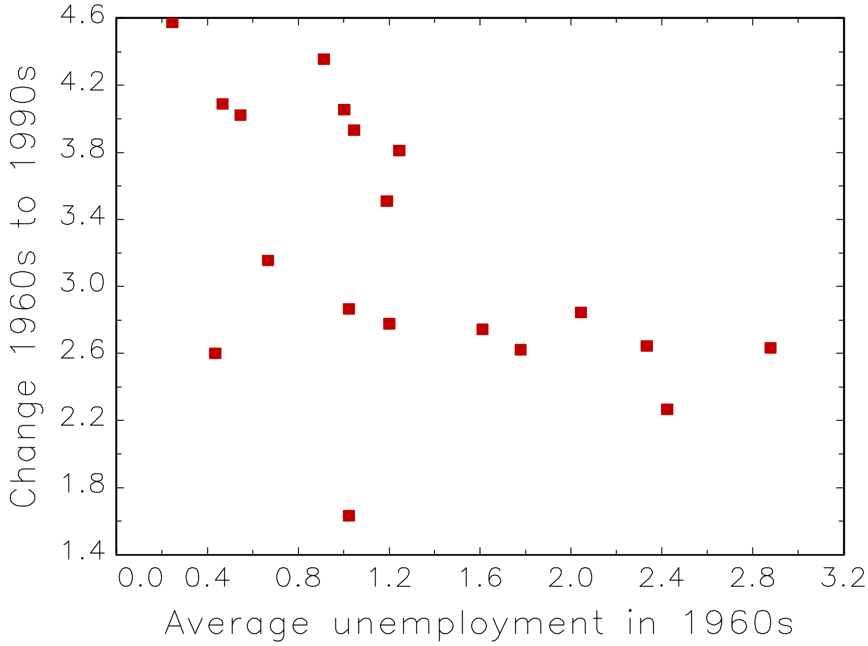

Figure 6: Cross plot of unemployment changes against initial levels.

Table 1: Augmented Dickey-Fuller Regressions

County	α_i (t_{α_i})	β_i (t_{β_i})	η_i (t_{η_i})	$-\eta_i/\alpha_i$	$\overline{R}^2$
1 Østfold	−0.125 (−2.16)	0.524 (3.49)	0.338 (1.84)	2.70	0.27
2 Akershus	−0.105 (−2.32)	0.675 (4.99)	0.121 (1.61)	1.15	0.42
3 Oslo	−0.081 (−2.74)[c]	0.798 (6.54)	0.123 (1.50)	1.52	0.55
4 Hedmark	−0.142 (−2.53)	0.570 (3.98)	0.352 (2.07)	2.48	0.32
5 Oppland	−0.143 (−2.47)	0.636 (4.46)	0.339 (2.07)	2.37	0.37
6 Buskerud	−0.117 (−2.34)	0.637 (4.53)	0.205 (1.83)	1.75	0.38
7 Vestfold	−0.086 (−1.59)	0.407 (2.44)	0.176 (1.27)	2.05	0.13
8 Telemark	−0.111 (−2.06)	0.529 (3.38)	0.301 (1.74)	2.71	0.26
9 Aust-Agder	−0.136 (−1.78)	0.218 (1.29)	0.400 (1.61)	2.94	0.06
10 Vest-Agder	−0.097 (−2.02)	0.570 (3.91)	0.228 (1.66)	2.35	0.31
11 Rogaland	−0.114 (−1.87)	0.461 (2.84)	0.222 (1.53)	1.94	0.18
12 Hordaland	−0.117 (−1.80)	0.384 (2.34)	0.300 (1.51)	2.56	0.13
13 Sogn og Fjordane	−0.235 (−2.26)	0.225 (1.33)	0.396 (2.01)	1.68	0.09
14 Møre og Romsdal	−0.161 (−2.06)	0.374 (2.30)	0.388 (1.78)	2.41	0.14
15 Sør-Trøndelag	−0.105 (−1.95)	0.524 (3.49)	0.288 (1.71)	2.74	0.26
16 Nord-Trøndelag	−0.123 (−1.74)	0.305 (1.79)	0.395 (1.64)	3.21	0.09
17 Nordland	−0.153 (−2.25)	0.473 (3.02)	0.531 (2.14)	3.47	0.22
18 Troms	−0.155 (−2.28)	0.529 (3.45)	0.490 (2.09)	3.16	0.26
19 Finnmark	−0.154 (−1.93)	0.338 (1.94)	0.580 (1.89)	3.76	0.10
Mean	−0.130 (−2.09)	0.483 (3.24)	0.325 (1.76)	2.47	0.24
Max abs value	0.235 (2.75)	0.798 (6.54)	0.580 (2.14)	3.76	0.55
Min abs value	0.081 (1.59)	0.218 (1.29)	0.121 (1.27)	1.15	0.06

Notes: Estimated coefficients with "t"-statistics in parentheses. Sample period is 1963–1998. Estimated equations are $\Delta U_{it} = \alpha_i U_{it-1} + \beta_i \Delta U_{it-1} + \eta_i + v_{it}$. Dickey–Fuller critical value: –3.62 (1%), –2.95 (5%), –2.61 (10%). Superscript a, b, and c denote rejection of the unit root null at 1%, 5% and 10% level, respectively.

Table 2: Panel data unit root tests (IPS)

Test statistic	Average value	5%[1]	1%[1]	$\Gamma_{\bar{t}}$ [2]	$\Gamma_{\overline{LM}}$ [2]
$t - bar$	$\bar{t} = -2.09$	–1.81	–1.98	–2.74	–
$LM - bar$	$\overline{LM} = 4.12$	3.76	4.17	–	2.14

Notes: The *t–bar* and *LM–bar* statistics are computed using the results in Table 1.

1) Exact sample critical values obtained from Im *et al.* (1997), Tables 3 and 4.

2) $\Gamma_{\bar{t}}$ and $\Gamma_{\overline{LM}}$, defined by equations (8) and (10) in the text, converges weakly to standard normal variates as both N and $T \to \infty$.

Table 3: The Additive Outlier Model

County	Break year	η_i (t_{η_i})	δ_i (t_{δ_i})	α_i (t_{α_i})
1 Østfold	1978	1.106 (3.91)	2.789 (7.17)	−0.339 $(-3.98)^c$
2 Akershus	1986	0.511 (3.48)	2.013 (7.71)	−0.344 $(-4.82)^b$
3 Oslo	1986	0.500 (2.33)	3.475 (9.12)	−0.296 $(-5.60)^a$
4 Hedmark	1986	1.846 (9.52)	1.996 (5.79)	−0.321 $(-3.94)^c$
5 Oppland	1985	1.832 (8.63)	1.876 (5.18)	−0.292 (−3.81)
6 Buskerud	1984	0.958 (4.50)	2.049 (5.84)	−0.269 $(-3.95)^c$
7 Vestfold	1982	0.514 (2.42)	3.086 (9.45)	−0.224 (−3.32)
8 Telemark	1985	1.600 (6.20)	2.892 (6.55)	−0.275 (−3.34)
9 Aust-Agder	1979	1.284 (5.58)	2.858 (8.78)	−0.569 $(-4.02)^c$
10 Vest-Agder	1978	0.678 (2.68)	2.907 (8.34)	−0.334 $(-4.43)^b$
11 Rogaland	1985	1.192 (6.92)	2.239 (7.60)	−0.389 (−3.86)
12 Hordaland	1980	1.005 (4.17)	3.023 (8.64)	−0.419 (−3.74)
13 Sogn og Fjordane	1979	0.921 (6.07)	1.437 (6.69)	−0.741 $(-5.10)^a$
14 Møre og Romsdal	1978	1.239 (5.54)	2.191 (7.10)	−0.479 $(-4.10)^c$
15 Sør-Trøndelag	1985	1.692 (7.34)	2.716 (6.89)	−0.307 (−3.75)
16 Nord-Trøndelag	1978	2.078 (10.15)	2.022 (7.17)	−0.357 (−3.31)
17 Nordland	1979	2.395 (13.47)	2.015 (8.02)	−0.475 $(-4.20)^c$
18 Troms	1979	2.290 (13.13)	1.858 (7.53)	−0.462 $(-4.21)^c$
19 Finnmark	1985	3.012 (17.24)	1.996 (6.68)	−0.456 $(-4.05)^c$
Mean				−0.425 (−4.08)

Notes: Estimated coefficients with "t"-statistics in parentheses. Sample period is 1963–1998. Location of break year determined by minimizing t_α. Critical values (approximate): –4.95 (1%), –4.25 (5%), –3.90 (10%).Superscript a, b, and c denote rejection of the unit root null at 1%, 5% and 10% level, respectively.